Accounting for the source of exchange rate movements: new evidence

Katie Farrant[*]

Gert Peersman[**]

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* International Economic Analysis Division, Bank of England, Threadneedle Street, London EC2R 8AH. E mail: katie.farrant@bankofengland.co.uk.

** Ghent University, Department of Financial Economics, W. Wilsonplein 5D, B-9000, Ghent. E mail: gert.peersman@ugent.be.

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Abstract

This paper analyses the role of the real exchange rate in a structural vector autoregression framework for the United Kingdom, euro area, Japan and Canada versus the United States. A new identification strategy is proposed building on sign restrictions. The results are compared to the benchmark conventional approach of Clarida and Gali based on long-run zero restrictions. Although the restrictions are derived from the same theoretical model, the results are strikingly different. In contrast to the benchmark model, an important role for nominal shocks in explaining real exchange rate fluctuations is found.

Key words: Exchange rates, vector autoregressions.

JEL classification: C32, E42, F31, F33.

Summary

Considerable research has previously been carried out to try to explain past movements in exchange rates. We examine this issue by estimating a structural vector autoregression, with sign restrictions, for the United Kingdom, the euro area, Japan and Canada versus the United States. The structural vector autoregression identifies not only demand, supply and monetary policy shocks, which may be important in explaining exchange rate movements, but also specific exchange rate shocks. These exchange rate shocks can be thought of primarily as movements in the exchange rate which are not explained by fundamentals. As far as we are aware, this is the first time that specific exchange rate shocks have been identified using sign restrictions, which is a much more general and less stringent approach than traditional identifying procedures.

We find that, while fundamentals have been important in explaining movements in exchange rates, there are also specific exchange rate shocks that have had a significant influence in determining exchange rate paths over time. This is in contrast to a number of other studies, which suggest that exchange rate movements can primarily be explained by demand shocks. Applying the traditional identifying strategy based on long-run restrictions to our data set, however, supports the findings of these other studies, suggesting that the identification strategy is important in determining the results.

1 Introduction

Most empirical studies that try to distinguish between the real and nominal sources of exchange rate movements have used structural vector autoregressions (SVARs) to analyse the relative importance of different shocks in explaining exchange rate fluctuations. However, they disagree in their results. The source of the disagreement seems to be the strategy that is used.[(1)] In a seminal paper, Clarida and Gali (1994) examine the importance of nominal shocks in explaining real exchange rate fluctuations. They use a long-run triangular identification scheme proposed by Blanchard and Quah (1989) and King *et al* (1991). The nominal shocks are identified by assuming that such shocks do not affect real variables, ie the real exchange rate or output, in the long run. Doing this, they find that demand shocks explain the majority of the variance in the real exchange rate. These results are confirmed by Funke (2000) for the United Kingdom versus the euro area. Chadha and Prasad (1997) apply the Clarida and Gali (1994) approach to the Japanese yen/US dollar exchange rate and also find that demand shocks play a crucial role in explaining fluctuations, although supply shocks are also important.[(2)] On the other hand, Artis and Ehrmann (2000) estimate structural VARs and identify monetary policy and exchange rate shocks using short-run zero restrictions. More specifically, they assume that all nominal shocks have no immediate effect on output. This study finds that the exchange rate seems mostly to reflect shocks originating in the foreign exchange market itself. Canzoneri *et al* (1996) reach a similar conclusion. They estimate VARs for a number of European countries and check whether the most important shocks in explaining the variance decomposition of output are also the most important in explaining exchange rate fluctuations. Supply shocks explain most of the movement in output but can hardly explain any variation in exchange rates.[(3)] Astley and Garratt (1998) examine movements in sterling bilateral exchange rates and find that demand shocks play an important role in determining exchange rate movements. Supply shocks also play some role, while the role of monetary shocks is limited. Overall, there is not yet a consensus on the issue.

A crucial aspect in the structural VAR literature is the identification strategy used. As already

(1) For an overview of the empirical evidence, see MacDonald (1998) or Artis and Ehrmann (2000).

(2) Other papers that have used the Clarida and Gali framework include Bhundia and Gottschalk (2003), who examine the bilateral South African rand/US dollar exchange rate and focus on the very sharp depreciation in 2001 Q4. The authors find that nominal disturbances explain by far most of the rand depreciation in 2001 Q4. Wang (2004) examines the Chinese real effective exchange rate and, by contrast, finds that relative real demand and supply shocks account for most of the variations in the real exchange rate during the estimation period. The study also finds that supply shocks are as important as nominal shocks in accounting for real exchange rate fluctuations.

(3) We should, however, be cautious in interpreting these results. If the exchange rate perfectly stabilises demand shocks, for instance, these will hardly explain output variation while having an important effect on the exchange rate.

mentioned, Clarida and Gali (1994) use a set of long-run zero restrictions. Their restrictions are based on a small open macro model in the spirit of Dornbusch (1976) and Obstfeld (1985). Long-run restrictions are often criticised in the literature. From an empirical point of view, Faust and Leeper (1997) show that substantial distortions in the estimations are possible due to small sample biases and measurement errors when using zero restrictions in the long run. Moreover, some equilibrium growth models (for instance many overlapping generations models) allow for permanent real effects of nominal shocks because they can affect the steady-state level of capital. The same is true for models based on hysteresis. Artis and Ehrmann (2000) also introduce short-run restrictions. Short-run restrictions are very stringent and could be misleading. There is no theoretical reason to justify a zero contemporaneous impact of nominal shocks on output, and it is inconsistent with a large class of general equilibrium models (Canova and Pina (1999)).

Faust (1998), Uhlig (1999) and Canova and De Nicoló (2002) use sign restrictions to identify monetary policy shocks as an alternative and find a more important role for policy shocks in explaining output fluctuations. The advantage of their approach is that it is not necessary to impose zero constraints on the contemporaneous impact matrix or on the long-run effects of the shocks. Instead, restrictions which are often used implicitly, consistent with the conventional view, are made more explicit. Peersman (2004) generalises this approach to a full set of shocks, ie supply, demand, nominal and oil price shocks, and compares the results with a conventional identification strategy. He also finds that the identification strategy plays a significant role in the results. Impulse responses based on traditional zero restrictions can be considered as a single solution of a whole distribution of possible responses that are consistent with the imposed sign constraints.[4] Peersman (2004) shows that a number of impulse responses based on zero restrictions are located in the tails of the distributions of all possible impulse responses. As such, this type of restriction can be very misleading, in particular when trying to draw conclusions about the effects of nominal shocks.

In this paper, we also apply the more recent sign restrictions approach to identify exchange rate shocks. To our knowledge, this is the first attempt to identify exchange rate shocks with this alternative strategy.[5] Specifically, we estimate respectively a three and four-variables VAR for the euro area, United Kingdom, Japan and Canada versus the United States for the sample period 1974

(4) Note that this is the case when the sign restrictions are imposed as $\leqslant 0$ or $\geqslant 0$.

(5) Faust and Rogers (2003) use sign restrictions to identify monetary policy shocks and analyse the impact on the exchange rate. They do not identify specific exchange rate shocks (or aggregate supply and demand shocks).

Q1-2002 Q4. The variables included in the basic VAR are output, prices and the real exchange rate. Supply, demand and nominal shocks are identified based on the same theoretical model of Clarida and Gali (1994). Instead of imposing long-run restrictions, we identify the shocks derived from the short-run properties of the model. In an extended four-variables VAR, we also add the nominal interest rate and disentangle monetary policy shocks and specific exchange rate shocks.

We find that nominal shocks have permanent effects. In particular, we find a permanent effect on the real exchange rate in some countries. Moreover, a substantial contemporaneous effect on the level of output is found. Since the conventional solution based on zero restrictions can be considered as one possible solution of the whole range of solutions obtained with sign restrictions, it is possible to locate the conventional solution in this distribution. For the real exchange rate, in particular, this seems to be quite often in the tails of the distribution. This illustrates that traditional zero restrictions on the contemporaneous impact and the long-run effects of the shocks can be misleading. This is also shown by the forecast error variance decompositions of the real exchange rate. They reveal the important role that nominal shocks play in explaining movements in exchange rates. The role is much more important than in Clarida and Gali's (1994) results. On the other hand, the role is significantly smaller than the results obtained in Artis and Ehrmann (2000). We also find that demand shocks and monetary policy shocks play a significant role in determining the path of the exchange rate and that supply shocks have little or no role in determining the exchange rate, either in the short run or the long run. We should note, however, that given that VARs are backward looking, it is possible that the exchange rate is reacting to *expected* demand and/or supply shocks, rather than the exchange rate itself being an important source of shocks, as the variance decompositions indicate.

The rest of the paper is structured as follows. In Section 2, we first describe the model of Clarida and Gali. We update their results for the countries under investigation in this paper and use them as a benchmark for our results. Our alternative approach based on sign restrictions is presented in Section 3. In Section 4, we extend the VAR model to check for robustness. Finally, Section 5 concludes.

2 The benchmark model of Clarida and Gali

In this section, we reproduce the results of Clarida and Gali (1994) for the United Kingdom, euro area, Japan and Canada for an updated sample period. We first give a brief overview of their theoretical model because we also use the model to justify our restrictions. They present the following stochastic two-country open macro model based on Obstfeld (1985) and Dornbusch (1976) with sticky prices:

$$
\begin{aligned}
y_t^d &= d_t + \eta q_t - \sigma \left[i_t - E_t \left(p_{t+1} - p_t \right) \right] && \textbf{(1)} \\
p_t &= (1-\theta) E_{t-1} p_t^e + \theta p_t^e && \textbf{(2)} \\
m_t^s - p_t &= y_t - \lambda i_t && \textbf{(3)} \\
i_t &= E_t \left(s_{t+1} - s_t \right) && \textbf{(4)}
\end{aligned}
$$

All variables except interest rates are in logs and represent home relative to foreign levels. Equation **(1)** is an open-economy IS equation where relative demand for output (y_t^d) depends upon a relative demand shock (d_t), is increasing in the real exchange rate ($q_t = s_t - p_t$) and decreasing in the real interest differential. Equation **(2)** is a price-setting equation where the relative price level in period t (p_t) is a weighted average of the expected market clearing price and the price that would actually clear the output market in period t (p_t^e). Equation **(3)** is a standard LM equation where relative real money balances are positively related to relative output and negatively related to the relative level of the interest rate. The relative interest rate (i_t) in equation **(4)** is determined according to the interest parity condition where s_t is the nominal exchange rate.

Clarida and Gali (1994) introduce three stochastic shocks in this model – relative supply, demand and nominal shocks. y_t^s and m_t are simple random walks, while there is a transitory and a permanent component in the relative demand shock d_t:

$$
\begin{aligned}
y_t^s &= y_{t-1}^s + \varepsilon_t^s && \textbf{(5)} \\
d_t &= d_{t-1} + \varepsilon_t^d - \gamma \varepsilon_{t-1}^d && \textbf{(6)} \\
m_t &= m_{t-1} + \varepsilon_t^n && \textbf{(7)}
\end{aligned}
$$

This model can be solved for the long-run flexible-price rational expectations equilibrium and represented as follows:[6]

$$y_t^e = y_t^s \quad \textbf{(8)}$$

$$q_t^e = \left(y_t^s - d_t\right)/\eta + \left[\eta\left(\eta+\sigma\right)\right]^{-1}\sigma\gamma\,\varepsilon_t^d \quad \textbf{(9)}$$

$$p_t^e = m_t - y_t^s + \lambda\left(1+\lambda\right)^{-1}\left(\eta+\sigma\right)^{-1}\gamma\,\varepsilon_t^d \quad \textbf{(10)}$$

In the long run, relative output (y_t^e), the real exchange rate (q_t^e) and relative price levels (p_t^e) are driven by three shocks, supply (ε_t^s), demand (ε_t^d) and nominal (ε_t^n). Moreover, the system is triangular in the long run. Only supply shocks have an effect on the long-run level of relative output, supply and demand shocks influence the long-run level of the real exchange rate and all three shocks are expected to have an impact on relative prices in the long run. The latter restrictions are used by Clarida and Gali (1994) to estimate the model and identify the shocks. Specifically, they estimate a three variables VAR in first differences, $\Delta x_t = \left[\begin{array}{ccc} \Delta y_t & \Delta q_t & \Delta p_t \end{array}\right]$ with three structural disturbances, $\varepsilon_t = \left[\begin{array}{ccc} \varepsilon_t^s & \varepsilon_t^d & \varepsilon_t^n \end{array}\right]$. Nominal shocks are identified by assuming that such shocks do not affect the real exchange rate and relative output in the long run. On the other hand, relative demand shocks do not have an impact on long-run relative output.

As a benchmark for our results in Section 3 and Section 4, we update their results for the sample period 1974 Q1 - 2002 Q4. We consider the United Kingdom, euro area, Japan and Canada versus the United States.[7] Impulse response functions together with 84th and 16th percentiles error bands based on Monte Carlo integration are shown in Figure 1. These are generally consistent with the results of Clarida and Gali (1994). A positive supply shock has a positive effect on relative output and a negative effect on the relative price level in all countries under investigation. The impact on the real exchange rate is, however, different. The supply shock leads to a significant depreciation in the United Kingdom and Japan and a significant appreciation in the euro

(6) For full derivation of the model, see Clarida and Gali (1994).

(7) Data are from the Bank of England database. The euro-area data are a weighted average of the individual countries' data. Lag length is determined by standard likelihood ratio tests and AIC information criterium. The deterministic components are a constant and a trend. All variables in the four countries we study are stationary in first differences. The model of Clarida and Gali (1994) implies that there are no cointegration relationships among the levels. We could not reject the null of no cointegration among the levels for the United Kingdom, Japan and Canada using the procedure of Johansen and Juselius in CATS. For the euro area, however, we do find a single cointegrating vector. However, these tests have low power. Moreover, this is not the focus of this paper: we are trying to compare our results with those of Clarida and Gali (1994) to see whether the identification strategy matters.

area, while the effect in Canada is insignificant. The appreciation in the euro area is also found in Detken *et al* (2002). This rather perverse effect, often also found for other currencies, may be explained by the fact that such a shock is accompanied by an upward shift in the aggregate demand curve as a result of real wealth effects and a home bias in consumption (Detken *et al* (2002)).

A positive demand shock increases relative output temporarily, results in a significant real exchange rate appreciation and raises relative prices. The latter, however, is never significant in the long run. After a positive nominal shock, there is a temporary significant positive effect on relative output and a permanent increase in the relative price level. The real exchange rate depreciates significantly in the short run in the euro area, Japan and Canada, but the depreciation is insignificant in the United Kingdom. Generally, apart from the impact of supply shocks on the exchange rate, all impulse responses are consistent with the model developed by Clarida and Gali (1994).

In order to analyse what drives movements in exchange rates, Clarida and Gali (1994) decompose the variance in the real exchange rate from all three shocks. The results for our variance decompositions are reported in the first row of all panels in Table A, at the one-quarter, one-year and five-year horizons respectively. We find that relative demand shocks in all four countries play the dominant role. Specifically, relative demand shocks explain 84%, 75%, 80% and 89% of the variance of the real exchange rate in the United Kingdom, euro area, Japan and Canada respectively. Relative supply and nominal shocks are shown to play little role. This finding is consistent with Clarida and Gali's results for the United Kingdom and Canada. However, they find an important role for nominal shocks in Japan, which we do not find. This difference might be due to our longer sample period. Overall, we find strong evidence for the exchange rate being driven by real shocks when we use a triangular long-run identification strategy.

3 A simple model with sign restrictions

The identification strategy used in structural VARs can have an important influence on the results. Clarida and Gali (1994) use a set of long-run zero restrictions. The latter, however, are often criticised in the literature. For instance, Faust and Leeper (1997) show that substantial distortions might arise due to small sample biases and measurement errors when zero restrictions in the long run are used. In addition, some equilibrium growth models or models with hysteresis effects allow for permanent real effects of nominal shocks. Other studies, such as Artis and Ehrmann (2000) also introduce short-run restrictions to identify nominal shocks. Specifically, they assume a zero contemporaneous impact of such shocks on the level of output. There is no theoretical reason to justify this zero contemporaneous impact, and it is inconsistent with a large class of general equilibrium models.[(8)]

In this section, we use sign restrictions to identify the shocks to evaluate the robustness of the benchmark results. This means that different shocks are identified according to the direction of their impact on the variables in the system. No zero restrictions are necessary. Sign restrictions are introduced by Faust (1998), Uhlig (1999) and Canova and De Nicoló (2002) to identify monetary policy shocks and are generalised by Peersman (2004) to a full set of shocks. In addition, Peersman (2004) compares the results based on conventional and sign restrictions. If the conventional decomposition produces impulse responses which are consistent with the imposed sign conditions, the former can be considered as a single solution of all possible responses based on sign restrictions. Peersman (2004) shows that a number of impulse responses obtained with zero restrictions can be situated in the tails of the distributions of all possible impulse responses that are consistent with the expected signs. We follow a similar approach with respect to exchange rates.

The sign restrictions that we impose are generally accepted and based on the short-run dynamics of the Clarida and Gali (1994) model. Specifically, in a world of sluggish price adjustment $(0 < \theta < 1)$, the variables deviate from their long-run equilibrium as follows:[(9)]

$$y_t = y_t^e + (\eta + \sigma)\, \upsilon\, (1 - \theta) \left(\varepsilon_t^n - \varepsilon_t^s + \alpha \gamma\, \varepsilon_t^d\right) \tag{11}$$

$$q_t = q_t^e + \upsilon\, (1 - \theta) \left(\varepsilon_t^n - \varepsilon_t^s + \alpha \gamma\, \varepsilon_t^d\right) \tag{12}$$

$$p_t = p_t^e - (1 - \theta) \left(\varepsilon_t^n - \varepsilon_t^s + \alpha \gamma\, \varepsilon_t^d\right) \tag{13}$$

where $\alpha = \lambda\, (1 + \lambda)^{-1}\, (\eta + \sigma)^{-1}$ and $\upsilon = (1 + \lambda)\, (\lambda + \sigma + \eta)^{-1}$. A nominal shock boosts

(8) Artis and Ehrmann (2000) also estimate the VAR in levels instead of relative variables, which implies that they do not have to assume that the dynamics of the system are similar across the two countries under consideration, a problem that we and Clarida and Gali (1994) have. On the other hand, estimating the VAR in levels can lead to a substantial bias in the results, in particular when there is an important role for symmetric shocks across countries. The existence of the latter will result in a more important role for pure exchange rate shocks when the VAR is estimated in levels.
(9) See Clarida and Gali (1994).

relative output in the short run with sluggish price adjustment. In addition, the relative price level rises and there is a depreciation of the real exchange rate. Home relative to foreign output increases in the short run as a result of a demand shock, there is a positive effect on relative prices and an appreciation of the real exchange rate. After a supply shock, relative output rises and there is a fall in relative prices. According to the model, the real exchange rate should depreciate in the long run. The short-run effect is, however, uncertain and will depend on the magnitudes of η, θ, λ and σ. We can summarise the signs (in the case of a positive shock) of the impulse response functions from the model as follows:(10) (11)

	y	p	q
supply	$\frac{\partial y_t}{\partial \varepsilon_t^s} \geqslant 0$	$\frac{\partial p_t}{\partial \varepsilon_t^s} \leqslant 0$	$\frac{\partial q_t}{\partial \varepsilon_t^s}$?
demand	$\frac{\partial y_t}{\partial \varepsilon_t^d} \geqslant 0$	$\frac{\partial p_t}{\partial \varepsilon_t^d} \geqslant 0$	$\frac{\partial q_t}{\partial \varepsilon_t^d} \leqslant 0$
nominal	$\frac{\partial y_t}{\partial \varepsilon_t^n} \geqslant 0$	$\frac{\partial p_t}{\partial \varepsilon_t^n} \geqslant 0$	$\frac{\partial q_t}{\partial \varepsilon_t^n} \geqslant 0$

These restrictions are sufficient to uniquely disentangle all the shocks based on sign restrictions. The restrictions are also consistent with a large class of other theoretical models, in particular given the monetary policy strategy in the countries under investigation. A restriction on the response of the real exchange rate to a relative supply shock is not necessary. While a depreciation is expected in the long run, the short-run effect is uncertain in the Clarida and Gali (1994) model. Moreover, a positive supply shock may be accompanied by an upward shift in the aggregate demand curve if there is a rise in domestic real wealth and consumers have a home bias in consumption.(12) The data will determine the sign of this response.

We estimate exactly the same VARs as in Section 2, but use the above mentioned constraints to identify the shocks. A sign restriction on the impulse response of variable p at lag k to a shock in q at time t is of the form:

$$R_{j,t+k}^{pq} \gtrless 0 \qquad \textbf{(14)}$$

For output and prices, we choose a value of k, ie the time period over which the sign restriction is binding, being equal to four quarters. For the real exchange rate, we only impose a value of

(10) If we assume that $0 < \gamma < 1$, ie that the demand disturbance is partially reversed in the future.
(11) Note that these restrictions are imposed on the *level* of the responses, while the VARs are estimated in first differences.
(12) See Detken *et al* (2002) or Bayoumi and Eichengreen (1994).

$k = 1$.[13] Following Uhlig (1999), we use a Bayesian approach for estimation and inference which is consistent with the method used in Section 2. Because we use $\gtreqless$ restrictions, there is not one single solution of impulse response functions for each draw from the posterior, but a whole distribution of possible solutions.[14] We therefore take a joint draw from the posterior for the usual unrestricted posterior for the VAR parameters as well as a possible decomposition for this draw obtained from a uniform distribution. We then construct impulse response functions. If all the imposed conditions are satisfied, we keep the draw, otherwise the draw receives zero prior weight and is rejected. Based on the draws kept, we calculate statistics and report the median responses, together with 84th and 16th percentiles error bands. Full details of the methodology and implementation of the restrictions can be found in Peersman (2004).

Impulse response functions to a supply, demand and nominal shock are presented in Figures 2a - 2d for the United Kingdom, euro area, Japan and Canada respectively. Full black lines are the median of the posterior distribution, dashed black lines the 84th and 16th percentiles error bands. In addition, we present the point estimation (grey line) of the benchmark Clarida and Gali (1994) identification strategy from Section 2. The latter solution is in most cases consistent with our imposed sign conditions and is part of the posterior distribution.[15] As a consequence, we can situate the solution with conventional restrictions in the whole distribution of possible solutions based on sign restrictions.

A number of the impulse response functions are very similar, ie the point estimate based on traditional long-run restrictions is close to the median response of all possible solutions consistent with the imposed sign conditions or, at least, lies within the 16th and 84th percentiles. There are, however, some striking discrepancies. We focus on the response of the real exchange rate. The reaction to a relative supply shock is very comparable across both methods. The point estimate based on conventional restrictions always lies between the confidence bands of sign restrictions. We find an insignificant effect in the United Kingdom and Canada. There is a significant depreciation in Japan and a significant appreciation in the euro area. The result for Japan is consistent with the findings of Chadha and Prasad (1997). The finding for the euro area, also found by Detken *et al* (2002), is often called a perverse supply-side effect (MacDonald (1998)).

(13) Changing these values has no effect on the general robustness and conclusions of our results.
(14) All possible decompositions can be generated using a search algorithm as explained in Peersman (2004).
(15) The only exceptions are the responses of relative prices to a demand shock in the United Kingdom and Japan. In the former country, the point estimate is slightly negative at lag 0 and lag 2, but insignificant. For the latter, the point estimate falls (insignificantly) below zero after one lag.

The impact of a relative demand shock on the real exchange rate is similar in Canada for both approaches, but is substantially smaller in the United Kingdom, euro area and Japan with sign restrictions. The Clarida and Gali (1994) solution lies very close to the 84th percentile of all possible solutions in the United Kingdom and euro area. The effect lies even outside the error bands in Japan. This smaller appreciation with sign restrictions might explain the somewhat larger effect on relative output that we find using this method.

The most striking difference, however, is the impact of a nominal shock. Clarida and Gali (1994) impose the restriction that there is no long-run effect of a nominal shock on the real exchange rate. When we only impose the restriction that there is no appreciation in the short run, we find a substantial permanent effect of nominal shocks in all countries.[(16)] The magnitude of the impact is very comparable with the response to a relative demand shock. The conventional solution based on a zero impact in the long run always lies outside the error bands in three out of the four countries. This means that the traditional approach results in rather extreme solutions and might indicate that the restriction of a zero impact in the long run is too stringent.

This different result is also reflected in the variance decompositions reported in Table A. The second row in all panels shows the contribution of all shocks to the variance of the real exchange rate for the three-variables VAR with sign restrictions. On the one hand, the contribution of relative demand shocks is much lower when we use sign restrictions to identify the shocks. While demand shocks explain the majority of real exchange rate fluctuations in our benchmark model of Clarida and Gali, the role is much more subdued with our alternative approach. For instance, demand shocks explain 82% of the variance at a five-year horizon in the euro area (error bands of 66% and 92%) with conventional zero long-run restrictions. The relative importance falls, however, to 30% when we use sign restrictions (confidence bands of 8% and 58%). A similar picture emerges in all other countries. The only exception is Canada, where we only find a small reduction in the relative importance of demand shocks. On the other hand, we now find a substantial role for nominal shocks in explaining real exchange rate fluctuations. In particular, we find that nominal shocks explain 50%, 57%, 62% and 26% of the immediate (one quarter) movements in the United Kingdom, euro area, Japan and Canada exchange rates while this was only 4%, 11%, 8% and 5% respectively in our benchmark model. Even at the five-year horizon, we still find a contribution from nominal shocks in explaining real exchange rate movements of

(16) The long-run effect in Canada is, however, not significantly different from zero.

31%, 42%, 57% and 11% in the United Kingdom, euro area, Japan and Canada respectively, which is always significantly above zero. Even the upper error band of the conventional solution is always below the lower error band of all possible solutions consistent with sign restrictions.

Overall, we find a very important role for nominal shocks in explaining exchange rate fluctuations, at least in the United Kingdom, euro area and Japan. This finding contrasts with the conventional results of Clarida and Gali (1994). This finding is rather surprising because the only difference between the two approaches is that we use the short-run dynamics of the model to identify the shocks instead of the long-run properties of the same model. In the next section, we analyse the robustness of our results when we make a distinction between monetary policy shocks and pure exchange rate shocks.

4 An extended model with sign restrictions

Building on the empirical model outlined in Section 3, we now extend the VAR to four variables in order to check the robustness of our results. A lot of studies make a distinction between monetary policy shocks and pure exchange rate shocks.[17] Historically, part of exchange rate fluctuations might be explained as a reaction to relative monetary policy shocks. In the context of optimal currency areas, it might be relevant to exclude these from our analysis about the role of the exchange rate. We can easily do this by extending the basic VAR to four variables, $\Delta x_t = \begin{bmatrix} \Delta y_t & \Delta p_t & s_t & \Delta q_t \end{bmatrix}$ with four structural disturbances, $\varepsilon_t = \begin{bmatrix} \varepsilon_t^s & \varepsilon_t^d & \varepsilon_t^m & \varepsilon_t^q \end{bmatrix}$. In contrast with the basic model, we now include the interest rate differential, i_t, and make a distinction between relative monetary policy shocks (ε_t^m) and pure exchange rate shocks (ε_t^q).[18] The latter could be the result of a time-varying risk premium in the exchange rate or movements in the exchange rate that are not explained by fundamentals. Note, however, that the exchange rate shock could be purely real as the model includes the real exchange rate. In order to uniquely disentangle all four shocks, we have to add some additional restrictions. These commonly accepted restrictions are based on a typical aggregate supply and aggregate demand diagram, which remains the core of many macroeconomic textbooks. In addition, the restrictions are very plausible given the monetary policy strategy in the countries under investigation. First, we impose the restriction that the interest rate differential does not fall after a relative demand shock. Second, the interest rate differential does not fall after an exogenous depreciation of the exchange

(17) For example Artis and Ehrmann (2000).
(18) The interest rate differential enters in levels terms.

rate. Both movements are consistent with central bank behaviour in the face of shocks that have an effect on inflation. Finally, we identify a monetary policy shock as a shock which has the opposite sign on the interest rate differential with respect to relative output, relative prices and the real exchange rate. This means that a restrictive monetary policy shock, measured as a rise in the interest rate, does not lead to a rise in output and prices. In addition, there is no depreciation of the real exchange rate. All these restrictions are consistent with the literature on the monetary transmission mechanism, and can be summarised in the following matrix:

	y	p	i	q
supply	$\geqslant 0$	$\leqslant 0$	?	?
demand	$\geqslant 0$	$\geqslant 0$	$\geqslant 0$	$\leqslant 0$
monetary	$\geqslant 0$	$\geqslant 0$	$\leqslant 0$	$\geqslant 0$
exchange	$\geqslant 0$	$\geqslant 0$	$\geqslant 0$	$\geqslant 0$

Again, these restrictions are sufficient to identify all shocks. A restriction on the reaction of the interest rate differential and the real exchange rate to a supply shock is not necessary. Impulse response functions are shown in Figures 3a - 3d. The reaction to supply and demand shocks are very similar to the results in the three-variables VAR with sign restrictions, which we will not discuss further. After an expansionary monetary policy shock, we find a permanent effect on the real exchange rate. This effect is significant in all countries except Canada. In most countries, we also find a permanent effect on relative output. The magnitude of a pure exchange rate shock is substantial in the short run. In the long run, we find a persistent effect on the real exchange rate for the median of the posterior in all countries under investigation, but this effect is not significant given the width of the confidence bands. In addition, we do not find a significant permanent effect on relative output performance in the United Kingdom, euro area, Japan and Canada. However, we do find a substantial contemporaneous effect of both pure exchange rate shocks and monetary policy shocks on relative output. This is a confirmation of the results of Canova and De Nicoló (2002) and Peersman (2004) for monetary policy shocks, but now also found for exchange rate shocks. We can conclude that restricting the immediate impact of both shocks to be equal to zero might be too stringent and may bias the results and conclusions.[(19)]

These results are confirmed by the variance decompositions of the real exchange rates. The third row in each panel of Table A shows the contribution of all shocks based on the extended VAR.

(19) In many VAR models, for instance the model of Artis and Ehrmann (2000) and Peersman and Smets (2001), the exchange rate is ordered after output and prices, which can underestimate the impact of exchange rate shocks.

We still find that a much smaller proportion of movements are explained by relative demand shocks compared to the benchmark Clarida and Gali (1994) results. The estimated contribution is very similar to the results of the three-variables VAR. The contribution of monetary policy shocks varies across countries. For example, at the five-year horizon, the contribution of monetary policy shocks in explaining real exchange rate fluctuations goes from being very small in Canada (4%) to being relatively high in Japan (37%). The role of monetary policy shocks in explaining movements in the United Kingdom and euro-area exchange rates lies in between these two extremes (10% and 12% respectively). Pure exchange rate shocks explain a substantial part of real exchange rate fluctuations in the very short run. In particular, at a one-quarter horizon, this is 40%, 25%, 30% and 20% for the United Kingdom, euro area, Japan and Canada respectively. Beyond a five-year horizon, the role of pure exchange rate shocks is more limited: 8%, 10%, 12% and 5% in the United Kingdom, euro area, Japan and Canada respectively. Overall, we find an important role for monetary policy and pure exchange rate shocks. In particular, we find a substantial impact of the latter in the short run, although the long-run effect is rather subdued.[20]

5 Conclusions

In this paper, we have analysed the source of real exchange rate movements in a structural vector autoregression framework for the United Kingdom, euro area, Japan and Canada versus the United States. We have first reproduced the benchmark results of Clarida and Gali (1994) for our sample period and countries under investigation. Their long-run identification restrictions are based on a stochastic two-country open macro model with sticky prices. We find that most of the variation in real exchange rates can be explained by relative aggregate demand shocks.

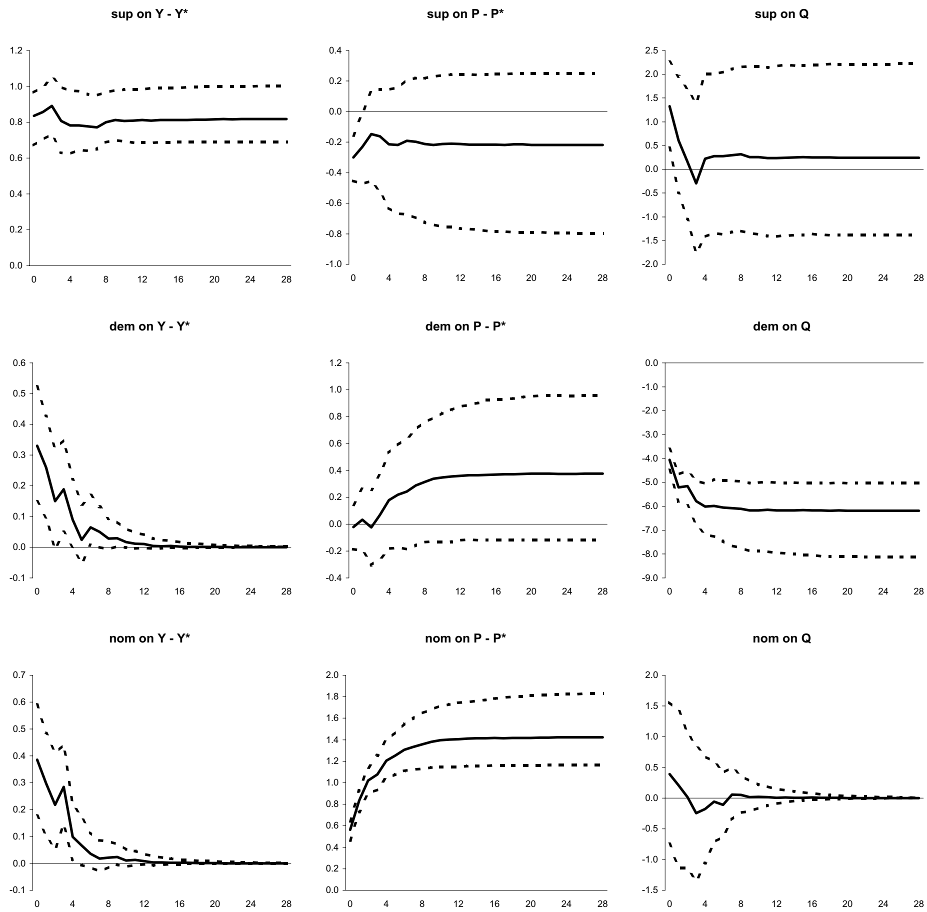
Long-run restrictions are, however, often criticised in the literature from a theoretical and empirical perspective. Using this type of restriction might be too stringent. We therefore introduce more recent and less restrictive sign restrictions as an alternative. The restrictions we implement are also based on the same theoretical model of Clarida and Gali (1994). Instead of focusing on the long-run properties of the model, we introduce restrictions based on the short-run dynamics of the model, which are also valid in a larger class of theoretical models. The advantage of our approach is that we have a whole range of possible impulse response functions, the classical

(20) Given that we find a very important role for nominal shocks in explaining exchange rate fluctuations in the three-variables model with sign restrictions and a rather limited role for monetary policy shocks in the four-variables model, we can suspect that the pure exchange rate shocks are predominantly nominal shocks.

solution based on zero long-run restrictions being one of them. This enables us to situate the classical solution in the whole distribution. We now find a notable important role for nominal shocks in explaining exchange rate fluctuations. The conventional solution based on long-run zero restrictions can be situated in the tails of the distribution of all possible solutions consistent with the short-run properties of the model. Even if we extend the model, and make a distinction between monetary policy shocks and pure exchange rate shocks, we still find an important role for the latter.

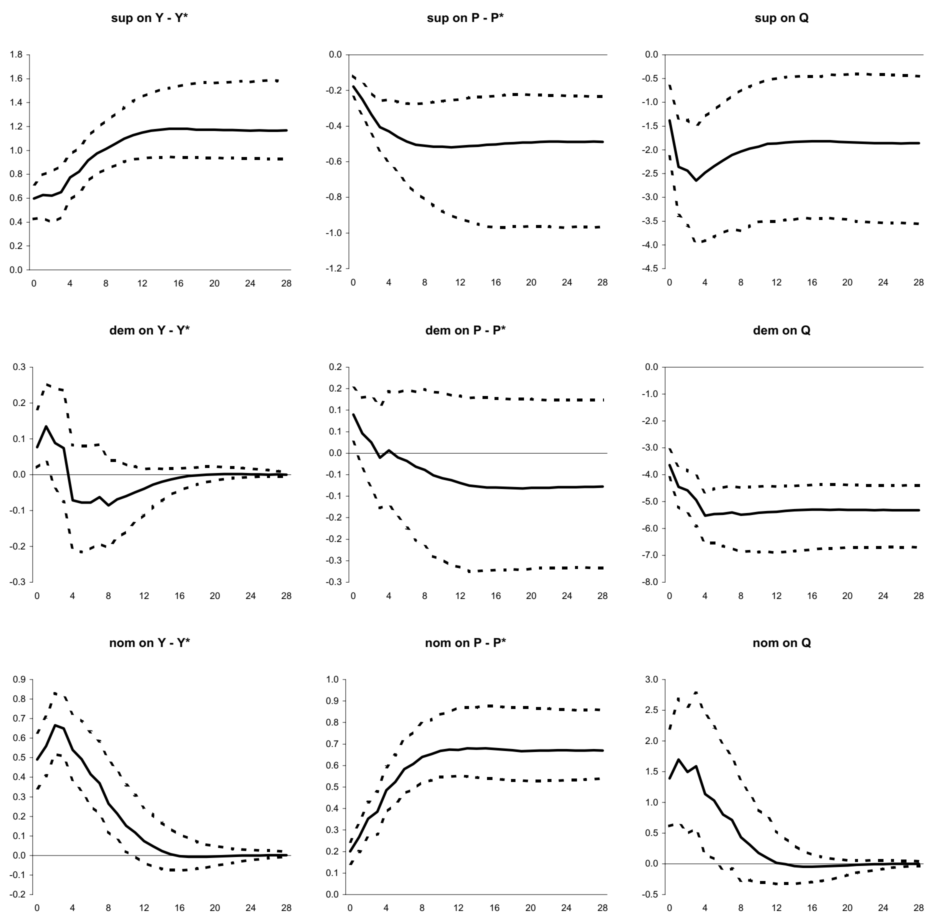
Focusing on the immediate impact, we find a substantial contemporaneous effect of both monetary policy and pure exchange rate shocks. The impact of the latter two is often restricted to be zero, and then used to identify the shocks. Our results indicate that this restriction might be too stringent and may bias the results.

Figure 1a - United Kingdom - Impulse responses based on Clarida-Gali identification



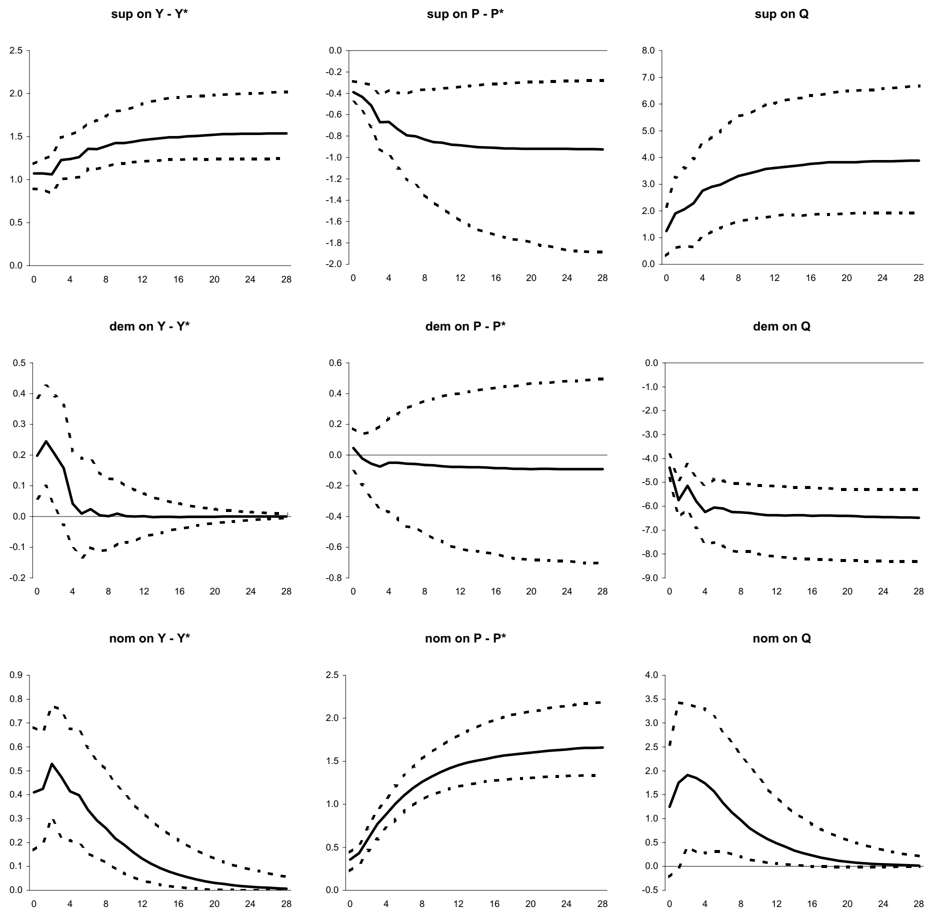
Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Figure 1b - Euro area - Impulse responses based on Clarida-Gali identification



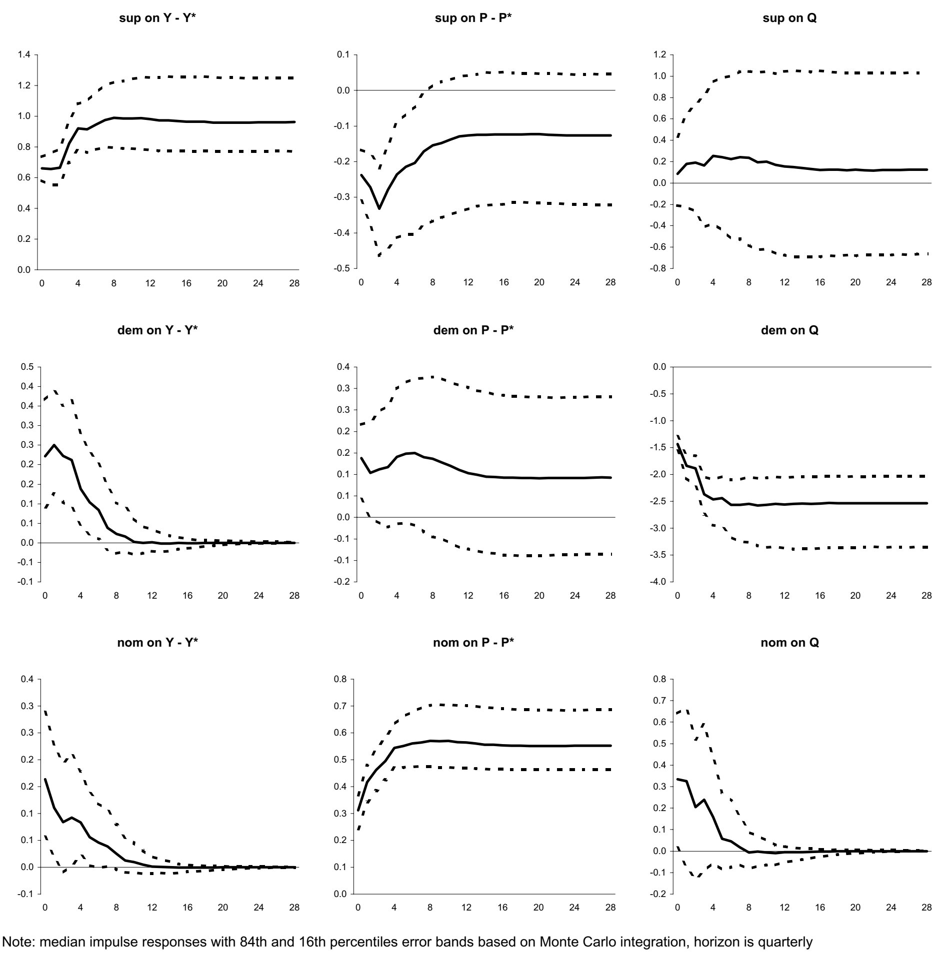
Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Figure 1c - Japan - Impulse responses based on Clarida-Gali identification



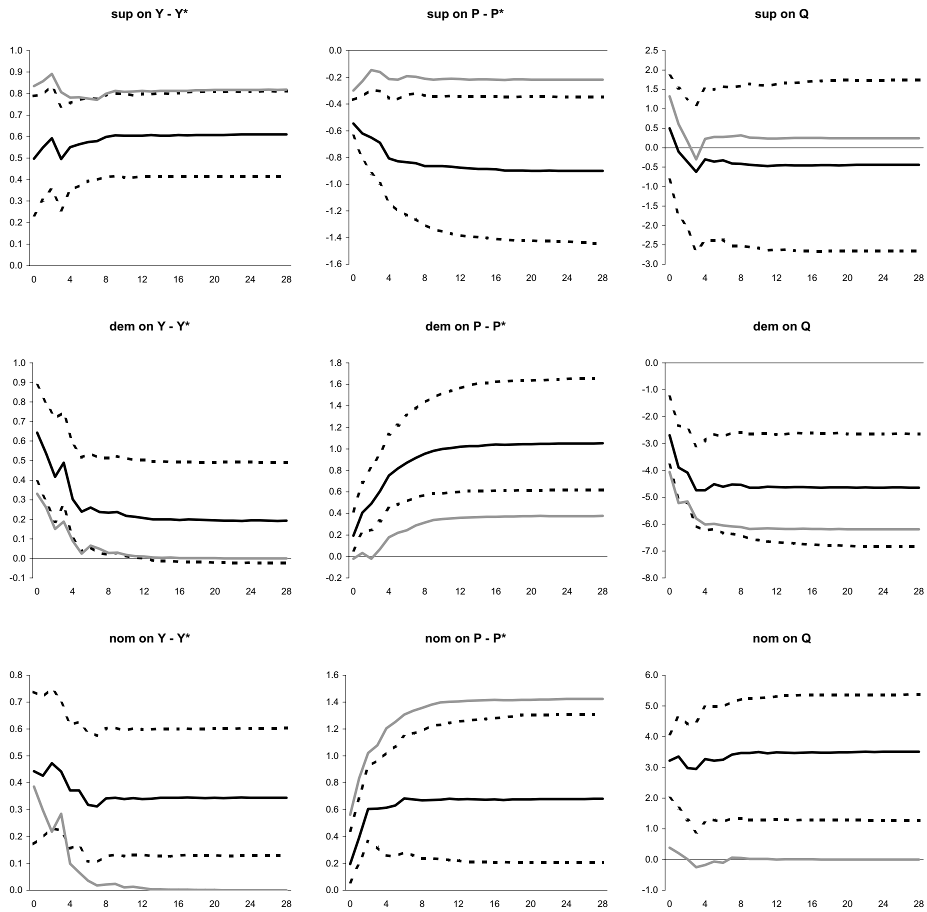
Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Figure 1d - Canada - Impulse responses based on Clarida-Gali identification



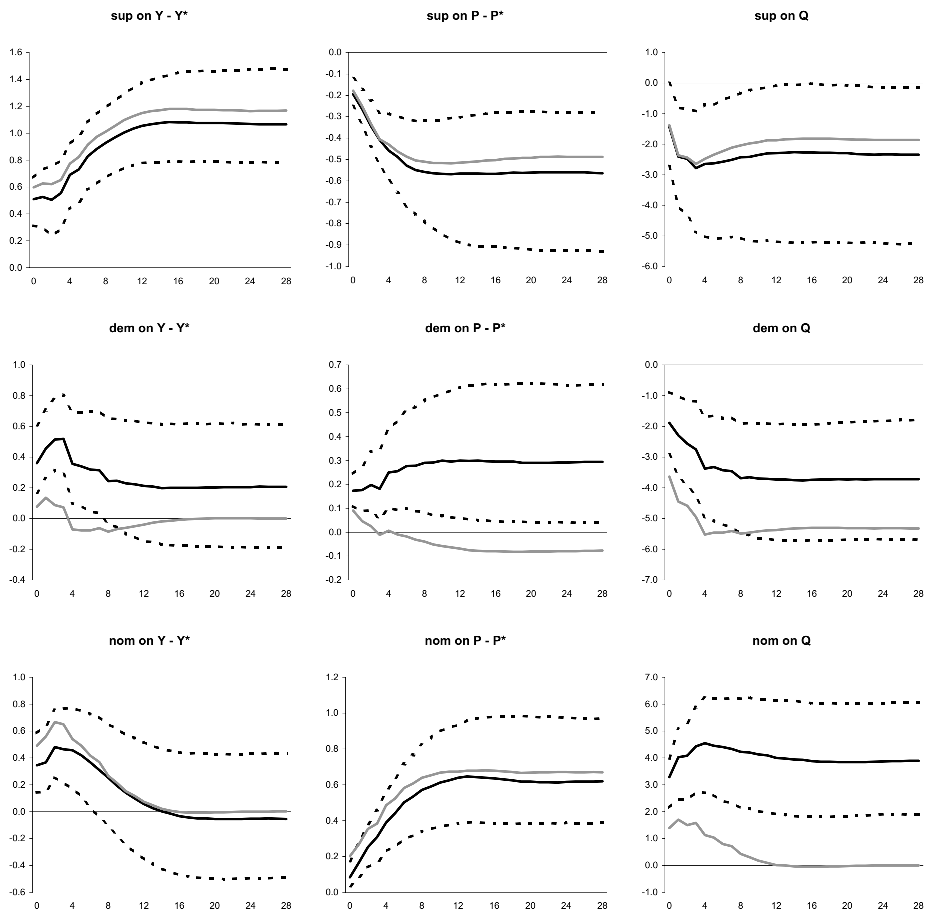
Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Figure 2a - United Kingdom - Impulse responses based on sign restrictions (3 variables)



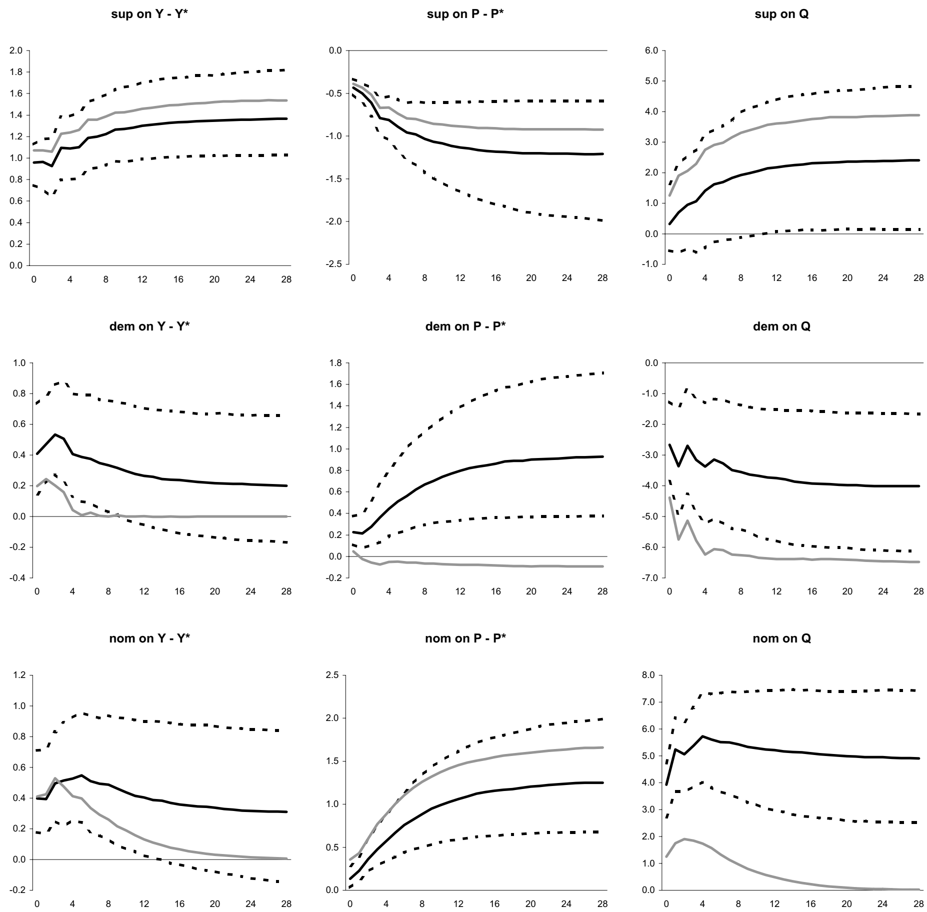
Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Figure 2b -Euro area - Impulse responses based on sign restrictions (3 variables)



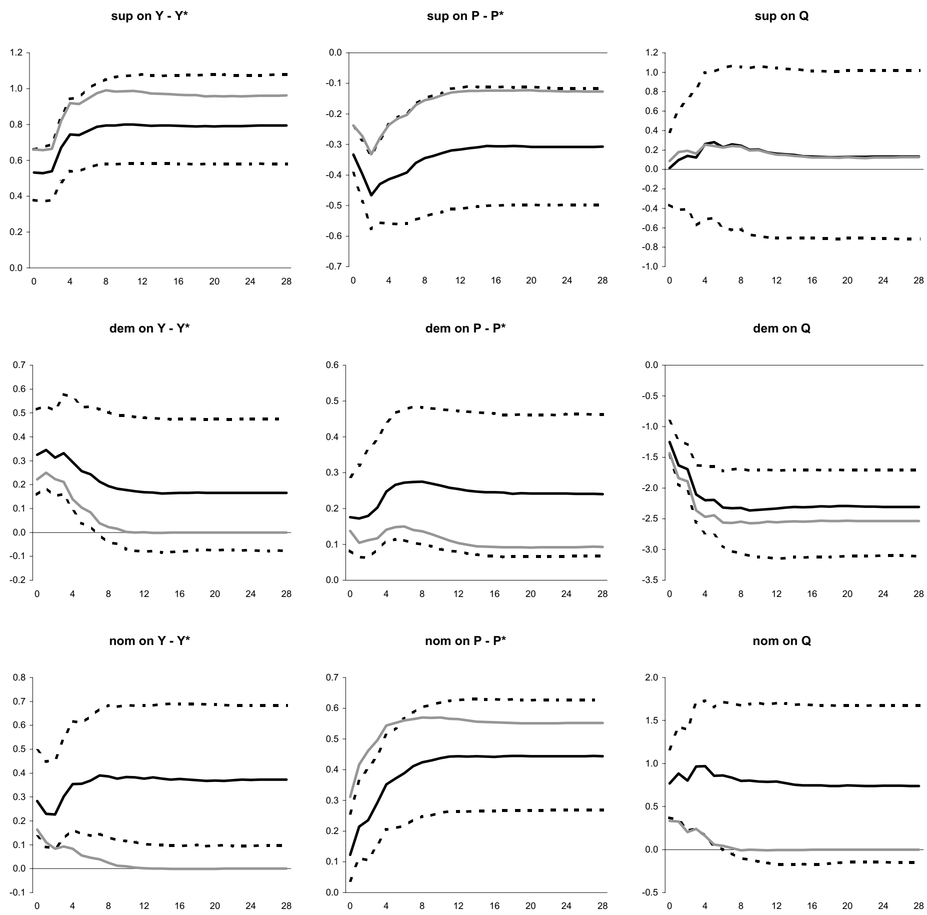
Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Figure 2c - Japan - Impulse responses based on sign restrictions (3 variables)



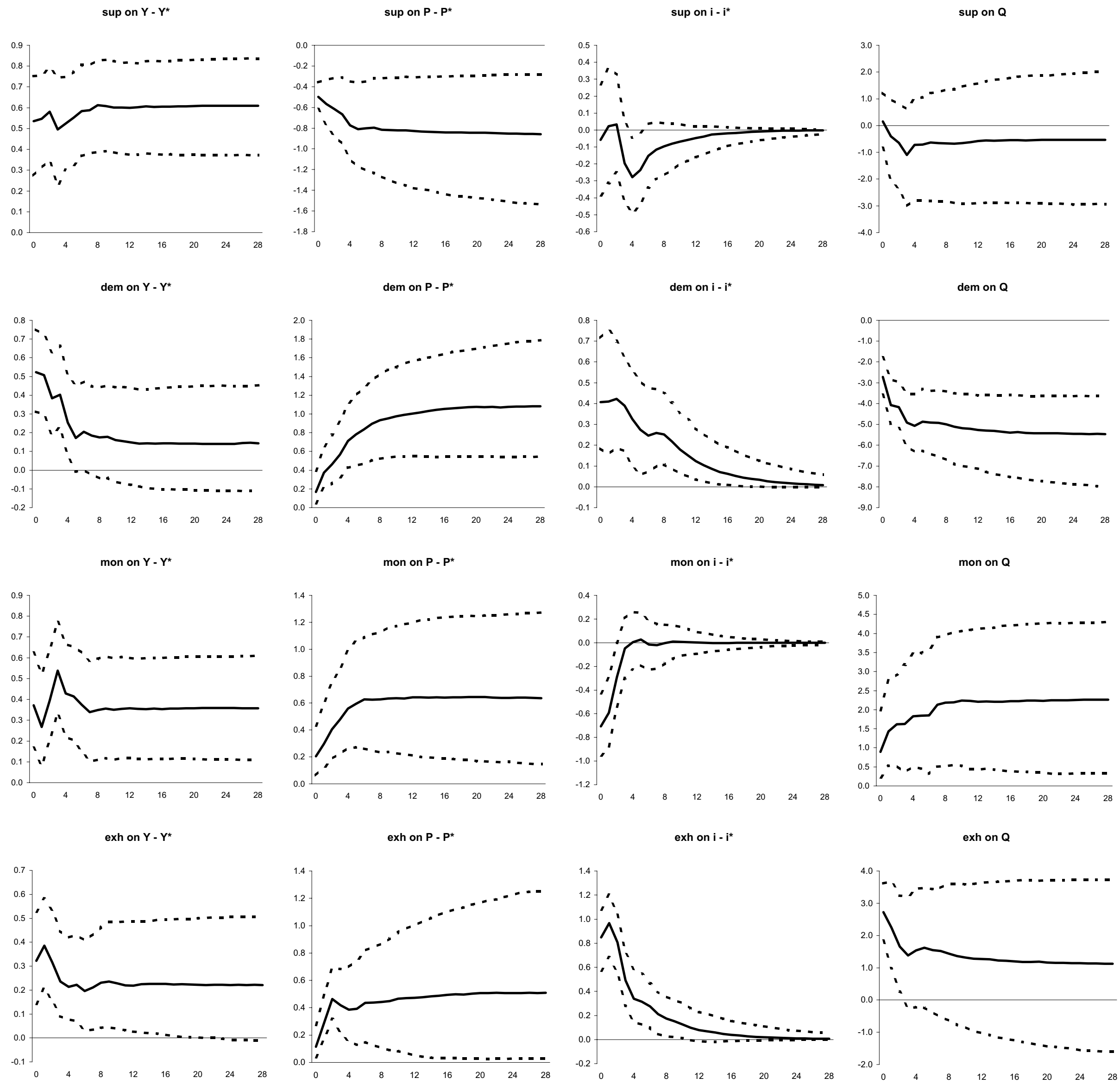
Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Figure 2d - Canada - Impulse responses based on sign restrictions (3 variables)



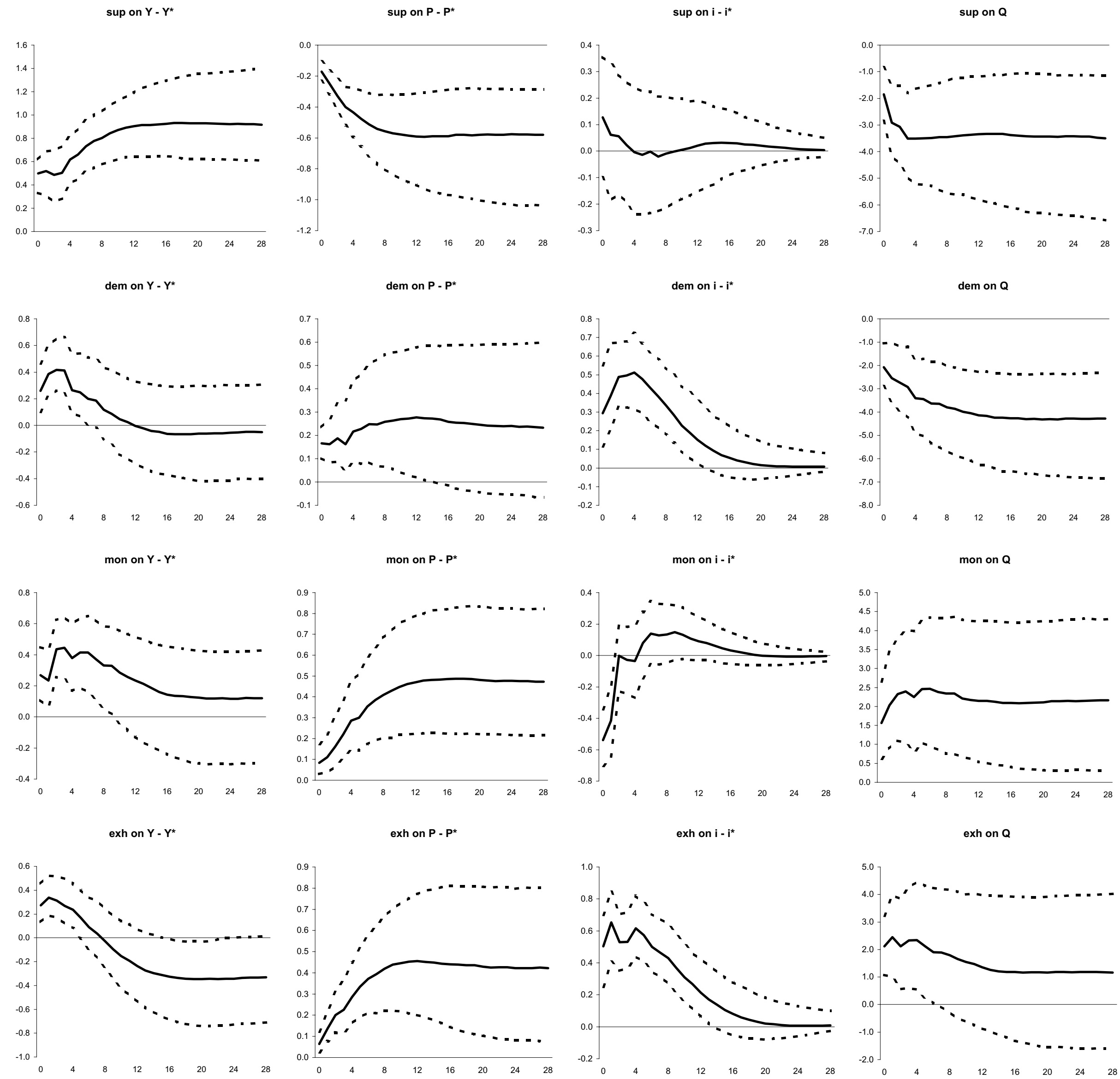
Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Figure 3a - United Kingdom - Impulse responses based on sign restrictions (4 variables)



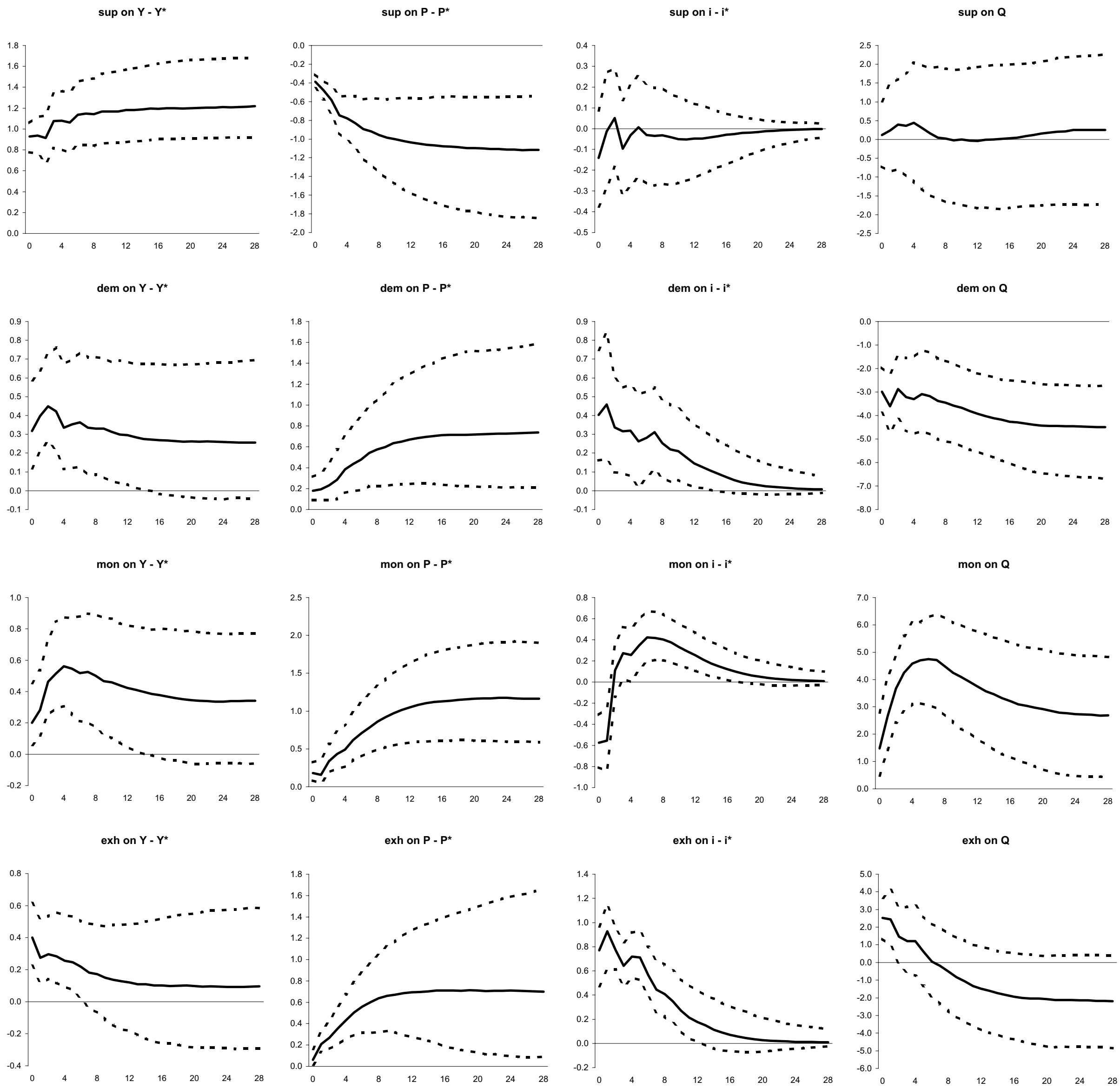
Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Figure 3b - Euro area - Impulse responses based on sign restrictions (4 variables)



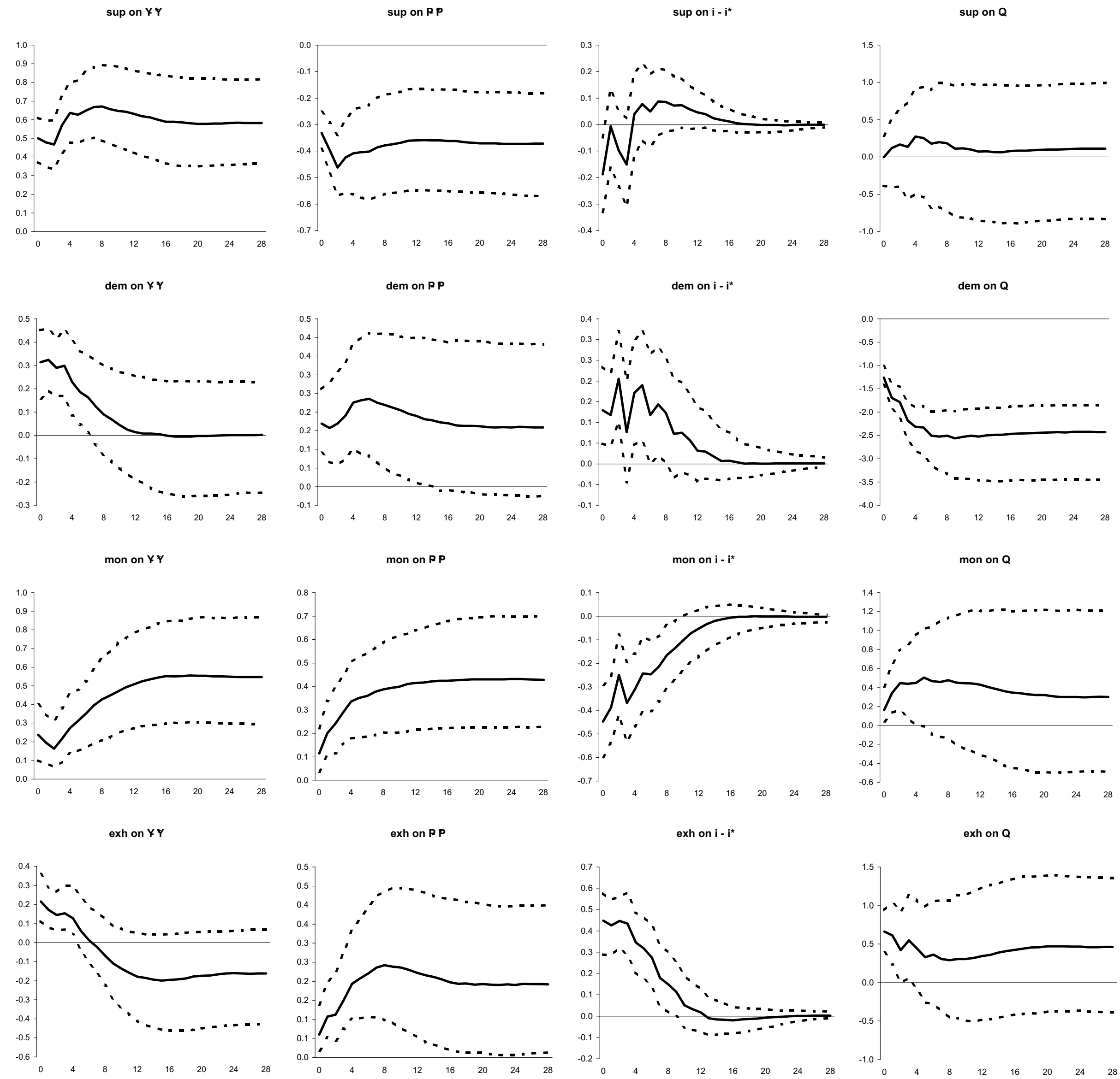
Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Figure 3c - Japan - Impulse responses based on sign restrictions (4 variables)



Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Figure 3d - Canada - Impulse responses based on sign restrictions (4 variables)



Note: median impulse responses with 84th and 16th percentiles error bands based on Monte Carlo integration, horizon is quarterly

Table A - Variance decompositions of real bilateral dollar exchange rates

United Kingdom

		supply			demand			nominal			monetary policy			exchange rate		
horizon	**method**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**
1 quarter	**Clarida - Gali**	0.08	0.25	0.01	0.84	0.95	0.66	0.04	0.14	0.00						
	Sign - 3 vars	0.04	0.22	0.00	0.37	0.73	0.07	0.50	0.81	0.20						
	Sign - 4 vars	0.03	0.14	0.00	0.40	0.69	0.17				0.04	0.20	0.00	0.40	0.70	0.18
1 year	**Clarida - Gali**	0.05	0.12	0.02	0.91	0.96	0.81	0.03	0.08	0.01						
	Sign - 3 vars	0.05	0.20	0.01	0.54	0.84	0.21	0.34	0.67	0.08						
	Sign - 4 vars	0.05	0.19	0.01	0.60	0.81	0.31				0.08	0.25	0.01	0.15	0.40	0.05
5 years	**Clarida - Gali**	0.04	0.14	0.01	0.95	0.99	0.86	0.01	0.02	0.00						
	Sign - 3 vars	0.05	0.22	0.01	0.56	0.86	0.21	0.31	0.66	0.06						
	Sign - 4 vars	0.05	0.21	0.01	0.62	0.83	0.32				0.10	0.32	0.01	0.08	0.30	0.02

Euro area

		supply			demand			nominal			monetary policy			exchange rate		
horizon	**method**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**
1 quarter	**Clarida - Gali**	0.11	0.25	0.03	0.75	0.91	0.55	0.11	0.26	0.02						
	Sign - 3 vars	0.12	0.37	0.01	0.19	0.47	0.05	0.57	0.87	0.25						
	Sign - 4 vars	0.19	0.44	0.04	0.24	0.47	0.06				0.13	0.38	0.02	0.25	0.56	0.07
1 year	**Clarida - Gali**	0.18	0.34	0.06	0.69	0.86	0.51	0.08	0.22	0.02						
	Sign - 3 vars	0.18	0.50	0.03	0.19	0.44	0.04	0.53	0.81	0.20						
	Sign - 4 vars	0.28	0.55	0.08	0.23	0.43	0.05				0.15	0.39	0.03	0.17	0.49	0.03
5 years	**Clarida - Gali**	0.14	0.30	0.04	0.82	0.92	0.66	0.02	0.07	0.01						
	Sign - 3 vars	0.16	0.47	0.03	0.30	0.58	0.08	0.42	0.73	0.13						
	Sign - 4 vars	0.26	0.56	0.06	0.34	0.60	0.11				0.12	0.34	0.02	0.10	0.36	0.02

Japan

		supply			demand			nominal			monetary policy			exchange rate		
horizon	**method**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**
1 quarter	**Clarida - Gali**	0.07	0.19	0.01	0.80	0.92	0.62	0.08	0.27	0.01						
	Sign - 3 vars	0.03	0.12	0.00	0.28	0.60	0.07	0.62	0.93	0.29						
	Sign - 4 vars	0.02	0.07	0.00	0.43	0.70	0.18				0.10	0.36	0.01	0.30	0.63	0.09
1 year	**Clarida - Gali**	0.10	0.25	0.02	0.75	0.88	0.58	0.09	0.27	0.01						
	Sign - 3 vars	0.03	0.16	0.01	0.24	0.55	0.04	0.67	0.91	0.35						
	Sign - 4 vars	0.03	0.09	0.01	0.34	0.60	0.13				0.35	0.62	0.15	0.14	0.41	0.04
5 years	**Clarida - Gali**	0.22	0.40	0.07	0.74	0.89	0.54	0.03	0.09	0.00						
	Sign - 3 vars	0.09	0.26	0.01	0.27	0.58	0.05	0.57	0.84	0.24						
	Sign - 4 vars	0.04	0.14	0.01	0.36	0.62	0.13				0.37	0.66	0.15	0.12	0.27	0.05

Canada

		supply			demand			nominal			monetary policy			exchange rate		
horizon	**method**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**	**median**	**upper**	**lower**
1 quarter	**Clarida - Gali**	0.02	0.09	0.00	0.89	0.97	0.75	0.05	0.18	0.00						
	Sign - 3 vars	0.03	0.13	0.00	0.69	0.92	0.35	0.26	0.58	0.06						
	Sign - 4 vars	0.02	0.10	0.00	0.71	0.88	0.48				0.01	0.07	0.00	0.20	0.40	0.07
1 year	**Clarida - Gali**	0.03	0.12	0.01	0.92	0.97	0.82	0.03	0.09	0.00						
	Sign - 3 vars	0.03	0.15	0.01	0.74	0.93	0.42	0.18	0.49	0.03						
	Sign - 4 vars	0.03	0.11	0.01	0.79	0.91	0.59				0.03	0.12	0.01	0.08	0.25	0.02
5 years	**Clarida - Gali**	0.05	0.16	0.01	0.94	0.99	0.83	0.00	0.01	0.00						
	Sign - 3 vars	0.05	0.18	0.01	0.79	0.94	0.49	0.11	0.38	0.02						
	Sign - 4 vars	0.05	0.14	0.01	0.78	0.91	0.58				0.04	0.16	0.01	0.05	0.18	0.01

Note: upper and lower bands are respectively 84th and 16th percentiles based on Monte Carlo integration

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